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The Dimensionality of Postmaterialism: An Application of Factor Analysis to Ranked Preference Data

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Previous research indicates that the measurement instrument for postmaterialist value-orientations is not one-dimensional, as Inglehart assumed. Instead, it seems to capture three independent dimensions suggesting an interpretation based on Habermas's theory of a 'colonization' of the lifeworld. In particular, the dimensions seem to catch different political potentials provoked by the penetration of the lifeworld by the econo-administrative complex. Two questions raised by these findings are addressed. First, they are based on a special measurement model for ranked preference data with largely unknown properties – thus, the model's ability to uncover the dimensional structure of Inglehart's instrument can be questioned. Second, the interpretation referring to Habermas requires that the dimensions are basically identical throughout advanced industrial countries and approximately stable over time. In this respect, the earlier findings are inconclusive. Results suggest that the measurement model accurately estimates number and structure of the latent dimensions implied by a set of rankings. In addition, the claim that Inglehart's instrument captures several independent and approximately identical dimensions in different industrial countries is confirmed with respect to Germany, the Netherlands, and the United States. Results also indicate that the dimensions are stable within countries.

Introduction

Sociological research on changing value orientations has been strongly influenced by Inglehart's concept of postmaterialism (Inglehart, 1977, 1989). Compared to other value-concepts, it has been more widely used over a relatively long time period. In addition, Inglehart's findings have attracted remarkable attention outside the scientific community. The scientific reception, as well as the public reception, was mainly focused on the controversial expectation of intergenerational value change towards a more postmaterialist society. This crucial assumption is supported by some newer studies, although the predicted cohort differences seem to be superimposed by period effects (Inglehart, 1989; de Graaf, 1989; Mnich, 1989; Klein, 1991; Sacchi, 1992; Abramson and Inglehart, 1995).

The ongoing debate about the stability of postmaterialist value orientations at both the cohort and the individual level has largely displaced the controversy about the dimensionality of postmaterialism. Several researchers concluded that the instrument designed to capture postmaterialist value orientations is not one-dimensional as Inglehart assumes. Their results rather seem to suggest that the instrument taps several value dimensions (van Deth, 1983*b*; Herz, 1979, 1987; Flanagan, 1982, 1987; Jagodzinski, 1984; Bean and Papadakis, 1994; Klein, 1995). Up to now, the dimensionality problem has remained unresolved, however, because the ranked preference data used for measurement have hindered a conclusive analysis. The main reason is that the standard methodology for solving dimensionality problems – factor analysis – should not be applied to rankings, since the evaluations of the individual items are not independent (Inglehart, 1979: 315). Therefore, most of the above-mentioned studies are based on a modified version of Inglehart's measurement instrument, which is suitable for factor analysis. Correspondingly, the findings are only valid regarding the modified version. Moreover, the modified versions are based on ratings, which are not very appropriate for the measurement of values, as I shall argue below. The few studies addressing the dimensionality problem on the basis of Inglehart's original instrument either ignore the fact that factor analysis should not be applied to rankings (Bean and Papadakis, 1994), or they use alternative statistical procedures, which do not allow unambiguous identification of the captured value dimensions (Herz, 1979; van Deth 1983*b*). In sum, research is inconclusive regarding the dimensionality of Inglehart's instrument, although there is evidence that the instrument is not one-dimensional.[1]

In a previous study, I have presented additional support for the notion that Inglehart's measurement instrument taps several value dimensions (Sacchi, 1994). The findings are based on a special factor model for ranked preference data, which was used to reanalyse some important studies on postmaterialism. Results suggest that Inglehart's original instrument taps three independent dimensions of political values or conflict. The dimensions are remarkably similar throughout Western industrial countries, and they appear approximately unchanged during the 1970s and 1980s. Thus, a common political cleavage structure of advanced industrial societies seems to be reflected by the three value dimensions. The proposed theoretical explanation for these findings builds on Habermas's theory of communicative action, or, more precisely, on the thesis of the 'colonization of the lifeworld' (Habermas, 1987, 1981). Basically, it is maintained that in developed capitalist societies the reproduction of the lifeworld is distorted by the expansion of media-driven subsystems, namely, the economy and the state. The colonization metaphor stands for this process characterized by the extensive monetarization and bureaucratization of society, which, ultimately, is responsible not only for symptoms of reification and cultural impoverishment, but also for the rise of different types of new potential for protest. Interestingly enough, the three-dimensional cleavage structure caught by the Inglehart items seems to match Habermas's theoretical expectations very well, each dimension representing a different type of reaction provoked by the colonization tendencies.

The summarized findings are inconclusive in several respects, however. First, they are obtained with a special measurement model for ranked preference data, which builds on specific assumptions that can barely be tested empirically. Also, little is known regarding the properties of the model, as it has been applied only rarely up to now. Therefore, its capacity to uncover the structure of the dimensions tapped by Inglehart's instrument remains rather unclear. Moreover, there are no unequivocal statistical criteria to determine the number of relevant dimensions in the context of the model. Thus, the results reported earlier may be questioned regarding both the number and the internal structure of the latent dimensions.

Second, possible objections can also be directed against the notion that Inglehart's instrument essentially captures identical and stable dimensions in most industrial countries. This conclusion has mainly been drawn from a visual comparison of loading patterns over countries and time. However, when the respective differences are not negligible, visual inspection opens a wide latitude for interpretation, and a systematic test regarding country variation and aggregate stability would be desirable. Obviously, both points are critical for the interpretation of the findings. Neither the claim that the items tap three independent dimensions nor their interpretation can be defended without the assumption that the measurement model performs well. In addition, the interpretation referring to Habermas requires that the dimensions are approximately equal in all industrial countries and that they remain more or less stable over time. This follows from the claim that they reflect a common cleavage structure emerging from general and long-term modernization processes.

In the present article, I hope to clarify the critical points listed above. In particular, I will address three related questions:

1. Does the measurement model for ranked preference data give accurate estimates of latent dimensions implied by a set of ranked items?
2. Do the Inglehart items tap three approximately identical, political value or conflict dimensions in different industrial countries?
3. Are the respective dimensions stable within countries?

In order to address these questions, two types of data will be analysed. On the one hand, artificial data with a known dimensional structure are analysed to assess the properties of the measurement model. On the other hand, data from the Political Action Panel, collected between 1974 and 1981 in Germany, the Netherlands, and the United States will be reanalysed (Zentralarchiv für empirische Sozialforschung, 1979, no date). In the first section, the application of the measurement model will be exemplified using data from the first panel wave. In the second section, I will investigate the performance of the measurement model by means of a Monte Carlo simulation. Thirdly, I will be concerned with the importance of national differences in the dimensions. Finally, the two panel waves will be analysed simultaneously, so that the stability of the dimensions on aggregate, as well as at the individual level, can be assessed.

Application of the Factorial Model for Ranked Preference Data

Usually, values are defined with reference to Kluckhohn (1976: 395) as a conception 'of the desirable which influences the selection from available modes, means, and ends of action.' Thus, in contrast to attitudes, values are seen as stable elements of personality with a strong impact on individual behaviour. Definitions of this type are relatively undisputed.

The appropriate way to measure values is far less consensual, however. None the less, there are 'strong theoretical arguments for choosing rankings instead of ratings when political values are the object of measurement.' (van Deth, 1983*b*: 410; Hellevik, 1994). Their theoretical superiority follows from the fact that ranking procedures force respondents to choose among items: respondents are most likely to reveal their values when goal conflicts and/or scarce resources make a choice between different claims inevitable. In contrast, rating methods – the most common alternative – allow respondents to give equal importance to every item. Ratings thus measure respondents' attitudes regarding the desirability of specific item-contents, rather than values. Moreover, since value-items in general are highly evaluated almost by definition, one cannot expect much variation in the answers. As a consequence, rankings are far better suited to the measurement of deep-rooted preferences, i.e. values, with high relevance for action. From this point of view, the ranking procedure applied by Inglehart contributes substantially to the strong impact of postmaterialist values on various forms of conventional and unconventional political participation (Barnes and Kaase, 1979; Inglehart, 1989, 1990; Jennings *et al.*, 1990; Sacchi, 1992). For similar reasons, the most important approaches to the study of value orientations do at least partly rely on ranked preference data.[2]

There are also possible objections to the application of ranking procedures. First, the criticism is made that respondents have no opportunity to give equal importance to different value-items and, consequently, such procedures imply that values are hierarchically ordered (Klages, 1992: 26; Böckler *et al.*, 1991; Maag, 1989, 1991). Now, as mentioned above, ranking procedures do indeed require that respondents' value orientations enable them to make a choice when faced by unresolvable goal-conflicts or by a measurement strategy. In this slightly weaker formulation, on the other hand, the assumption does not seem unreasonable, since respondents have to choose among conflicting claims not only in the interview situation, but also in everyday life. Moreover, in the case of the Inglehart items, this argument is corroborated by empirical evidence: measured by the rate of refusals, respondents do not seem to have special difficulties in ranking the items – an indication that the underlying assumption is appropriate (Zentralarchiv für empirische Sozialforschung, 1979: 44 f.).

A second and more serious objection to the application of rankings is the absence of a standard methodology for analysing ranked preference data. Factor analysis should not be used, since the scores (i.e. ranks) of the individual items sum up to a constant, and, therefore, are not independent (Jackson and Alwin, 1980). When there are only two items to be ranked, for example, it is obvious that there will be a perfect negative correlation between the resulting scores. Although the strength of the interdependence diminishes as the number of items increases, ranks remain negatively correlated by definition. These so-called 'ipsative' properties of ranked preference data violate a basic assumption of factor analysis, and, therefore, impose serious limitations on data analysis.

To overcome the second objection, Jackson and Alwin proposed a measurement model appropriate for factor analysis of ranked preference data (Jackson and Alwin, 1980; Alwin and Jackson, 1981).[3] Their model is built on the basic assumption that rankings reflect underlying evaluations, i.e. ratings, at an invariant scale, and that the ipsative properties stem solely from the ranking procedure. It is difficult to test this crucial assumption directly, since the supposed evaluations are not identical to observable ratings that allow respondents to change their evaluation standard from one item to another. Nevertheless, the few comparisons of ratings and rankings for the items applied by Inglehart to some degree support the notion that observed rankings reflect underlying ratings at constant scales. It turns out that the two methods only differ for a group of respondents giving high scores to all items when rating is applied – apart from that, rankings and ratings give similar results (van Deth, 1983*b*; Bacher, 1987).[4]

Starting from the observed rankings, the measurement model is designed to uncover the factor structure of the underlying non-ipsative and scale-invariant evaluations. The original Alwin–Jackson model assumes that respondents rank only one set of items. Inglehart's most elaborate measure of post-materialism, however, relies on two independent rankings of four and eight items, respectively (Table 1). This does not impose serious obstacles to analysis, because a slightly modified version of the measurement model can be applied to independently ranked items, as de Graaf *et al.* (1989) have shown.

Following Alwin and Jackson, two measures have to be taken to control for ipsativity in factor analysis. First, the rank of one item selected at random is excluded from analysis – the excluded information is redundant since it is implied in the rankings of the other items. In the case of several independent rankings, one item out of each has been excluded (de Graaf *et al.*, 1989). Second, for every item a phantom variable is introduced into the factor model, absorbing the negative correlations among the rankings induced by the ranking procedure. Figure 1 gives an example of a model with two common dimensions or factors in a LISREL-notation. The model is made up of two common factors (η_1, η_2), ten non-redundant rank indicators (y_1–y_{10}), and twelve phantom variables ($\eta_3 - \eta_{14}$) correcting for the

Table 1. *Item-sets of Inglehart's measurement instrument*

First ranking	
Item 1	Maintain order in the nation
Item 2	Give people more say in the decisions of the government*
Item 3	Fight rising prices
Item 4	Protect freedom of speech*
Second ranking	
Item 5	Maintain a high rate of economic growth
Item 6	Make sure that this country has strong defence forces
Item 7	Give people more say in how things are decided at work and in their communities*
Item 8	Try to make our cities and countryside more beautiful*
Item 9	Maintain a stable economy
Item 10	Fight against crime
Item 11	Move towards a friendlier, less impersonal society*
Item 12	Move towards a society where ideas are more important than money*

Note: Postmaterialist value-items are marked with asterisks.

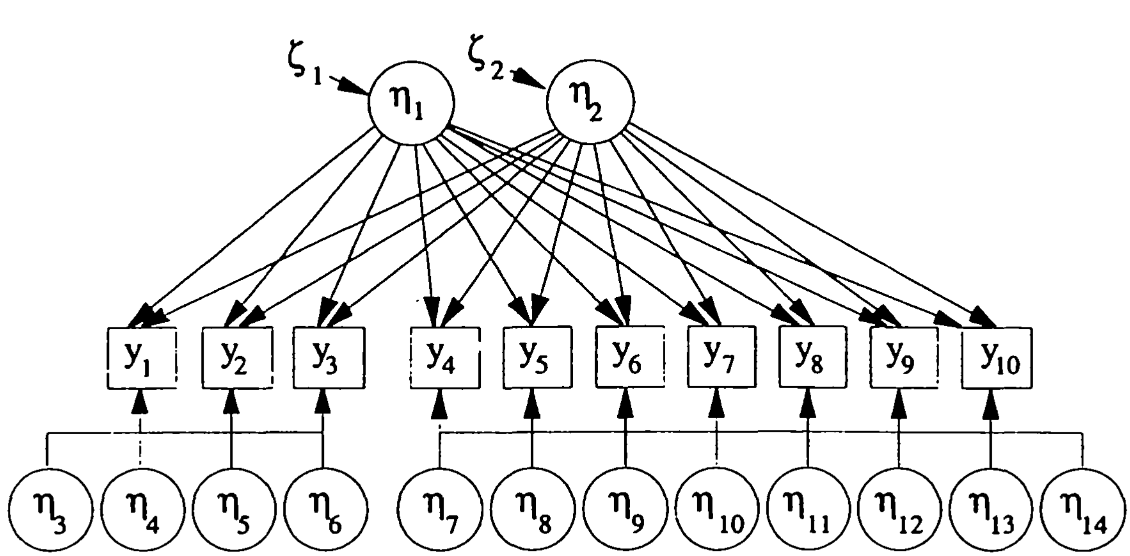

Figure 1. *Two-dimensional measurement model for postmaterialism.*

correlated error structure. Although the model contains no separate error terms, measurement error is taken into account by means of the phantom variable (Jackson and Alwin, 1980: 227).

Because measurement is based on two independent rankings in the case of the Inglehart items, the phantom variables or error variances are only correlated within each item set. The model also includes two phantom variables for the omitted rank-indicators (η_3, η_{14}), controlling their negative correlation with the analysed rankings (y_1–y_{10}). Theoretical arguments referring to indicators with ipsative properties shows that the correlations between phantom variable (η_i) and rank indicators (y_i) are a function of the number of ranked items. Specifically, the phantom variable η_i assigned to a particular rank-indicator y_i, which is part of a ranking of n items, is expected to exert a negative effect on the other rank-indicators, which amounts to $-1/n$. At the same time, the effect of the phantom variable on its 'own' indicator equals $1 - 1/n$.[5]

Following these general rules, the respective parameters of the λ_y-matrix can be fixed when the model is estimated, and, because the latent dimensions are assumed to be independent, the factor loadings are identified. A formal derivation for the special case with two independent rankings is found in de Graaf *et al.* (1989). It should be noted that the estimated factor loadings are deviations from the mean loading on the original non-ipsative factor. Since these deviations sum to zero for each ranking, the loadings of the excluded items can be calculated by adding all loadings of an item set and reversing the sign (de Graaf *et al.*, 1989: 187). Nevertheless, the resulting loading pattern can be interpreted as in conventional factor analysis, at least as long as the loadings of the original non-ipsative structure can be expected to have opposite signs.

The measurement model will first be applied to data from the first wave of the Political Action Panel, carried out in 1974 in Germany, the Netherlands, and the United States (Zentralarchiv für empirische Sozialforschung, 1979, no date). Despite considerable panel attrition, especially in Germany, the samples seem to be representative of the adult population (Hagenaars *et al.*, 1990). The analysis includes adult respondents at least 20 years old. Rank indicators are coded as a simple numeration according to subjective priority, where the most important item is coded as one.[6] The data-set and coding are thus identical to those used by de Graaf *et al.* in estimating the intragenerational stability of postmaterialist value orientations. Moreover, I have also excluded the rankings of the same items when estimating the model. Regarding the first factor, results can therefore by compared.

The parameters of the model are estimated with LISREL by the method of unweighted least squares (ULS) (Jöreskog and Sörbom, 1989). Ranked preference data in general, as well as in the present case, do not have a multinormal distribution, which is required for maximum likelihood estimation (ML) (Alwin and Jackson, 1981: 322). In contrast, the fit function of ULS can be justified without this requirement.

It is recommended to start the analysis with a model without common factors. This allows us to judge the improvement of fit associated with the extraction of the first factor (Jackson and Alwin, 1980: 228). After that, I will first estimate a model with one factor and then extend it step by step for additional dimensions. As in conventional factor analysis, this raises the problem of when the extraction has to be stopped. In principle, this could be decided with the χ^2-test for the respective LISREL-Model, which, however, would again require a multinormal distribution. Since this assumption is not met, there is no unequivocal criterion for deciding on the number of relevant dimensions (Jackson and Alwin, 1980: 228). For the moment, I assume that the χ^2-statistic is unbiased despite the skewed distributions of the rank indicators, and, consequently, that it can be used to decide on the number of dimensions. The performance of the χ^2-test for this type of model and data will be examined in detail in the next section.

The baseline model without a common dimension is expected to show a satisfactory fit, if the ranking procedure is the only source of correlation among the indicators. The results listed in Table 2 indicate that this is not the case. Compared to the 43 degrees of freedom, the χ^2-statistic of the baseline-model is very high in all countries.

The extraction of the first common value-dimension improves the fit considerably: there is a large χ^2-gain of several hundred points, compared to a loss of ten degrees freedom, and the goodness-of-fit index (GFI) indicates a better fit of the one-dimensional model as well (Table 2). Despite the improvement, the fit is not sufficient to accept the model and a second dimension is extracted. In order to identify the two-dimensional model, the loadings of the first factor have to be fixed on the estimates obtained from the one-dimensional solution (Jöreskog and Sörbom, 1986: III. 103). Since the factors are orthogonal, it is possible to extract them step by step. It should be noted that due to the fixed loadings, the sub-model estimated in the second step has ten degrees of freedom more than the complete two-dimensional model. It turns out that the extraction of the second dimension clearly improves the fit in all countries. The most substantial improvement is found in Germany, where the model with two factors can be accepted. For the other countries, the fit is still not satisfactory, and, consequently, a third factor is extracted. Again, the loadings of the first two factors are fixed, when the additional dimension is estimated. This extension of the model leads to an acceptable fit in the remaining countries as well. Compared to the degrees of freedom, the χ^2-gain is substantial, but smaller than for the previous factors.

The results of the χ^2-tests are further confirmed when the residuals are inspected. A large residual indicates that the respective model is not able to explain the respective correlation between two indicators. This may also occur when the overall fit is good. For the preferred models, however, residuals are small and residual plots indicate a close correspondence between model and data.[7]

Table 2. *Number of extracted factors and fit of the measurement model*

Number of factors	df[a]	West Germany (N = 822) χ^2	P	GFI	The Netherlands (N = 557) χ^2	P	GFI	United States (N = 780) χ^2	P	GFI
0	43	548.9	0.000	0.891	657.0	0.000	0.827	532.6	0.000	0.895
1	33	240.6	0.000	0.967	146.3	0.000	0.983	212.2	0.000	0.982
2	33 (23)	42.1	0.132	0.995	84.9	0.000	0.992	103.3	0.000	0.995
3	33 (13)	–	–	–	35.1	0.368	0.998	37.7	0.263	0.997

[a]Degrees of freedom for the respective extraction step and for the complete model (in brackets).

The superiority of the two- and three-dimensional models also becomes visible when the GFI are compared as an alternative measure of overall fit. Contrary to the measure reported up to now, the GFI is neither very sensitive to deviations from the normal distribution, nor does it depend on sample size. Unfortunately, the absolute level of fit cannot be judged with the GFI, since its statistical distribution is unknown (Jöreskog and Sörbom, 1986: I. 41). Hence, although the fit clearly improves as more factors are extracted, the significance of the gain cannot be assessed.

Disregarding the violation of the distributional assumptions for the moment, the conclusion seems clear: there are several dimensions underlying the ranking of the Inglehart items. This raises the question of how the dimensions have to be interpreted substantively. As in conventional factor analysis, one has to look at the estimated factor loadings to answer this question (Table 3). Unstandardized factor loadings are reported, since the standardized coefficients are difficult to interpret in the context of the model.[8]

Obviously, the first dimension corresponds to the well-known postmaterialist value dimension in all countries. As Inglehart would expect, the loadings of postmaterialist (marked with asterisks in the Table) and materialist items have opposite signs. In addition, country differences are small, although probably not random. When the loadings are compared with the estimates by de Graaf *et al.* (1989), the differences are negligible. With respect to the structure of the first value dimension, their findings are thus replicated by the present study. As a consequence, Inglehart's index construction is corroborated by the findings. However, his one-dimensional theory of value change cannot explain the additional dimensions, and therefore an alternative explanation is required. An alternative interpretation would be desirable too with respect to the first dimension, postmaterialism, since Inglehart's theory has a weak foundation, as many critics have noted (Thome, 1985; Flanagan, 1987; Böckler *et al.*, 1991; Klages, 1992).

As mentioned at the outset, the dimensions suggest an interpretation based on Habermas's (1987) theory of communicative action. The dimensions are then tied to different types of reactions provoked by the 'colonization' of the lifeworld, that is, its penetration by the growing econo-administrative complex. I have elaborated this interpretation elsewhere (Sacchi, 1994, in press). In the present context, I shall confine myself to a brief sketch of the basic idea. For this purpose, it is useful to succinctly recall Habermas's argument. Basically, he postulates that in modern societies, there is a close connection between growing monetarization and bureaucratization on the one hand, and aggravating symptoms of reification on the other hand. The tendencies to the former are pushed ahead by the dynamic of economic growth and the attempt to contain its negative consequences through political

Table 3. *Unstandardized factor loadings (λ_y)*

Items		Dimension 1 (Potential for resistance)			Dimension 2 (Potential for withdrawal)			Dimension 3 (Disorientation)	
(cf. Table 1)	No.	FRG	NL	USA	FRG	NL	USA	NL	USA
Maintain order	1	0.257	0.642	0.284	−0.552	−0.458	−0.465	0.076	0.180
More say: Government*	2	−0.294	−0.510	−0.210	0.438	0.372	0.496	−0.027	−0.211
Fight rising prices	3	0.253	0.169	0.292	0.029	0.173	0.007	−0.324	−0.245
Economic growth	5	0.663	0.512	0.746	0.680	0.424	0.253	0.033	0.351
Strong defence forces	6	0.649	0.828	0.851	−0.390	−0.204	−0.106	−0.086	0.125
More say: Work*	7	−0.046	−0.776	−0.488	0.749	0.586	0.956	−0.241	0.038
Stable economy	9	0.379	0.336	0.451	0.183	0.138	0.145	1.257	0.272
Fight against crime	10	0.590	1.150	0.459	−0.675	−0.162	−0.443	−0.385	−0.416
*Less impersonal society**	11	−0.944	−0.905	−0.978	−0.190	−0.456	−0.260	−0.061	−0.049
*Ideas versus money**	12	−1.145	−1.091	−1.006	−0.125	−0.455	−0.128	−0.320	0.412

Note: The items tied to the contradiction of system and lifeworld are in italics.

regulation and administrative intervention; thus, they are seen as inevitable concomitants of capitalist modernization. The tendency to reification refers to the consequences of the expansion of the econo-administrative complex from the perspective of the lifeworld.

The intervention of this complex into spheres which are crucial for the reproduction of the lifeworld – cultural reproduction, social integration, and socialization – implies reification and devaluation of cultural traditions. Essentially, the reason is that the steering media of the economic and political system – money and power – cannot substitute for communicative action in these respects. As a consequence, the intervention of the steering media in these spheres, that is the 'colonization', provokes various reactions on the side of the lifeworld. In the present context, one possible reaction is of special interest, namely, legitimation crises giving rise to potentials for protest.

Habermas (1987: ii. 578 f.; 1981) distinguishes two types of political potentials. The first corresponds to a conservative, anti-modernist potential, defending structural niches against the penetration by the steering media of economy and administrative system, the second represents a progressive, emancipatory potential that tries out new forms of co-operation and living together. The former 'potential for withdrawal' becomes manifest, for example, in the traditional middle-class protests against modern comprehensive schools or against rising taxes, whereas the latter 'potential for resistance' is mobilized by new social movements, particularly by alternative, ecology, peace, and youth movements.

Thus, the potentials represent specific types of opposition to the modernization values embodied in the institutional orders of the economy and the state. More specifically, they are both opposed to modernization, in so far as it implies economic growth and state expansion (Habermas, 1987: ii. 577). However, the two potentials advance their modernization critique from very different angles. In the case of the second, emancipatory potential the opposition is grounded in the anticipation of a more democratic and just social order. In contrast, the first potential is rooted in conservative or even authoritarian traditions appropriate for the defence of particularist interests and traditional forms of life. Although both potentials try to contain the further assimilation of the lifeworld into the expanding subsystems, their political alignment is very different: the former is attached to the 'new' right, the latter to the 'new' left.

With the above considerations in mind, one can now ask how members of both potentials are likely to rank the Inglehart items. To answer this question, it is useful to start with those items that apparently have an intimate connection to the opposition of system and lifeworld as described by Habermas. This holds true for three items. First, it has been mentioned that the critique of economic growth (item 5) is common to both potentials. Second, the item 'move towards a friendlier and less impersonal society' (item 11) points to disturbances in the process of social integration. Moreover, the opposition of 'friendlier' and 'less impersonal' implies that these disturbances are rooted in the substitution of communicative action by more anonymous forms of action coordination, such as Habermas's steering media. Third, the claim 'move towards a society where ideas count more than money' (item 12) essentially points to tensions between cultural values and economic constraints. In Habermas's terms, the item relates to distortions in the process of cultural reproduction due to an unbounded monetarization of society. It follows that members of both potentials should give low priority to economic growth, and at the same time have high priority for the containment of the pathological symptoms caused by the penetrating steering media, as expressed by items 11 and 12. In addition, the other items concerned with the economic system (items 3 and 9) are also likely to have rather low priority for both potentials. However, their evaluation of the remaining items can be expected to be clearly different. The left, emancipatory potential is most likely to combine its critique of modernization with claims for participation and democratization, as expressed by items 2 and 7. In contrast, the conservative potential combines a similar critique with insistence on the traditional social order to be defended against the disintegrating forces of modernization. Thus, confronted with the item-list, members of the conservative potential are likely to give high priority to items 1 and 10, expressing a strong concern for law and order. Hence, the positions of both potentials are antagonistic regarding the contradictory goals of further democratization versus the defence of the established order (see also Habermas, 1987: ii. 507).

The empirical findings meet these expectations very well, as the loading patterns in Table 3 indicate. In the case of the first dimension, a preference for democratization and the protection of the lifeworld against the penetrating steering media goes hand in hand with low priority for all economic issues in all countries. Thus, the postmaterialist pole of the factor is consistent with the expectation of an emancipatory potential reacting against the 'colonization' process. It is noteworthy that Habermas (1987: ii. 576 f.) posits explicitly this link between the rise of the political potentials reacting to the 'colonization' process and the growing share of postmaterialist values observed by Inglehart. Moreover, previous research demonstrates that postmaterialist values in fact are a strong determinant of protest activities and of membership in new social movements (Barnes and Kaase, 1979; Jennings *et al.*, 1990; Inglehart, 1990; Dalton and Kuechler, 1990; Opp, 1990 Rohrschneider, 1990; Sacchi, 1992, 1994). These findings corroborate the suggested notion of postmaterialism as a potential for protest.

For the second factor, the loading pattern is similar regarding the items directly pointing to the opposition of system and lifeworld. Hence, people expressing high priority for the protection of the lifeworld against the penetrating steering media are again grouped together at one pole. However, contrary to the postmaterialist potential captured by the first dimension, this group shows a strong preference for conservative claims for law and order (items 1 and 10), and strong defence forces (item 6). Thus, the factor perfectly meets the expectation of a conservative potential provoked by the 'colonization' tendencies. This interpretation of the loading patterns is also supported by previous studies on a much broader selection of industrial countries (Sacchi, 1994, in press). They reveal that the underlying potential corresponds to a conservative, religious group of predominantly elderly people. Most notably, this group expresses pronounced support for authoritarian and repressive policies, as advocated by the 'new' right.

Before interpreting factor three, it is useful to consider a possible objection to the outlined interpretation of dimensions one and two.[9] At a theoretical level, the emancipatory 'potentials for resistance' and the conservative 'potentials for withdrawal' represent mutually exclusive categories: a person cannot be modernist and anti-modernist at the same time. In contrast, the statistical model implies that there are at least some people with high scores on both dimensions, since the factors are orthogonal (i.e. uncorrelated). Accordingly, it could be concluded that the statistical model is incompatible with the theoretical assumptions. I think, however, that this conclusion is inappropriate for two reasons.

First, theory predicts that despite all ideological discrepancies both potentials are opposed to the further expansion of the econo-administrative complex. In particular, they share the critique of the negative consequences of economic growth. In this respect, a certain overlap of both potentials is evident, even though the remedies they advocate – democratization versus the defence of the established order – are mutually exclusive.

Second, even if one is ready to assume that the theory allows an unequivocal distinction between modernist and anti-modernist goals (i.e. items), it is an empirical question as to whether or not such logical contradictions are reflected by the evaluations of the respondents. From this point of view, the high internal consistency of the two main dimensions inherent in the rankings is the remarkable finding, and not the observation that some respondents express preferences, which are self-contradictory in theoretical terms. People's claims are quite often paradoxical; for example, members of ecology movements generally advocate leftist and participatory reforms, while at the same time being clearly anti-modernist in other respects (Christmann, 1992; Rucht, 1988: 307 f.). Taking the above arguments together, it is evident that an overlap of these two potentials does not undermine the suggested interpretation.

Turning to factor three, the loading patterns are difficult to interpret with reference to Habermas's theory. In addition, only some of the loadings are substantial and country differences seem more pronounced. Nevertheless, construct validation for a larger sample of industrial countries has revealed that the third dimension seems to be invariantly tied to an anomic, lower-class potential recruiting itself from the declining 'old' left (Sacchi, 1994, in press). Its outstanding characteristics are high scores on various indicators for political disorientation and alienation. Although such potential is not explicitly predicted by Habermas, it is in line with his argument. Specifically, he maintains that the interplay of colonization tendencies and the growing distance between expert cultures and the broader public implies disorientations and political apathy (Habermas, 1987: ii. 521 f.). It seems plausible that these phenomena are more prevalent among lower social strata with less education and cognitive resources. Hence, the type of potential tied to the third factor is consistent with his theoretical expectations.

Moreover, the loading patterns also gain a certain plausibility when interpreted against this backdrop. First, the loadings essentially indicate that the potential puts most weight on the items 'fight rising prices' and 'fight crime', while giving low priority to the items referring to the economic system. It seems reasonable that the former goals have a more obvious usefulness for disoriented people with a poor understanding of politics, than the relatively abstract claims referring to the performance of the economic system. Second, if the underlying potential is interpreted as a specific reaction to 'colonization' tendencies, it has to be explained why the contradiction of system and lifeworld is not consistently reflected by the loadings of the respective items (5, 11, and 12). Some of these loadings are close to zero, while the loading of item 12, which is substantial, does not show the expected sign for the United States. However, if disorientation is in fact the main characteristic of the underlying potential, its recognition of the contradiction between system and lifeworld will not be consistent almost by definition. Correspondingly, the loadings of the respective items are almost accidental and difficult to interpret. In sum, the loading pattern is quite understandable when related to the aforementioned construct validation, which shows that the third factor captures a lower-class potential of disorientated and alienated people.

Whether or not the suggested interpretation of the multi-dimensional solution is valid strongly depends on the assumption that the dimension are accurately estimated by the applied measurement model. Moreover, any interpretation of the above results is contingent on this assumption. In the next section I will therefore investigate the properties of the measurement model, especially its ability to assess number and structure of latent dimension implied by a set of rankings.

The Monte Carlo Simulation

Monte Carlo experiments can be defined as a method of investigating the performance of statistical models and estimation techniques by applying them to artificially generated data-sets (Vogt, 1993). Hence, Monte Carlo simulation is an obvious way to examine some properties of the measurement model for ranked preference data, which has rarely been applied up to now. Thus, to some degree its applicability remains unclear (Jackson and Alwin, 1980: 229).

Below, the properties of the model will be tested with artificial data-sets that are drawn from the observed ranking data. The construction of data-sets for simulation relies on various theoretical assumptions about the structure of observed data. Because the structure of the artificial data is known, it can be used to test the ability of the model to uncover its 'true' structure. A limitation of this research strategy, however, is that the results cannot be generalized, given the specific assumptions introduced for simulation. For the present article, the Monte Carlo experiment is designed to answer the following questions:

1. Is there a reliable criterion, and if so, which one, for deciding on the number of factors, that is, the dimensionality, within the framework of the Alwin–Jackson model for ranking data?
2. Do the loading patterns of the factors properly reflect the structure of the underlying dimensions; and is the measurement model in this respect really superior to conventional factor analysis?

To start with, artificial data with known dimensionality, structure, and properties similar to those of the observed rankings have to be generated. In order to create such data, four assumptions about the measurement process and the nature of the latent value dimensions have to be introduced. Regarding the value dimensions, I first assume that they are independent and normally distributed. Second, it is assumed that the items have clear and identical meanings for all respondents, implying that they unambiguously mark positions in a value space defined by the independent dimensions. This is rather a strong assumption, of course, which, however, will be somewhat relaxed later by introducing an error term. Third, it is taken for granted that the evaluation of the items reflects the degree of agreement between item content and respondents' value orientations. Consequently, these evaluations may be expressed as (Euclidean) distances in the value space:

$$d_{ij} = \sqrt{\sum_{k=1}^{n} (f_{ki} - f_{kj})^2 + \varepsilon}$$

where

d_{ij} Euclidean distance between item i and respondent j, including measurement error
f_{ki} position of item i on value dimension k
f_{kj} position of respondent j on value dimension k
n number of value dimensions
ε measurement error

Hence, the smaller the Euclidean distance between respondent i and the position of item j, the 'closer' respondents' values and item-content. The Euclidean distances thus correspond to scale-invariant ratings of the items. As required by the measurement model, these evaluations – which unfortunately cannot be measured directly in empirical studies – do not have ipsative properties until they are subject to ranking (Jackson and Alwin, 1980: 222). From the third assumption, it follows that respondents, when asked to rank the items, will give highest priority to the item with the smallest d_{ij}, second priority to the item which is next, and so on. The transformation of the Euclidean distances into rankings is thus straightforward. The fourth and last assumption refers to measurement error. It is supposed that respondents' evaluations contain an uncorrelated, normally distributed error component (ε), representing, for example, random disturbances from the interview situation, or hazy item-formulations. The error is expected to affect only respondents' evaluations (d_{ij}), not the ranking procedure itself. This seems to be a reasonable assumption, however, as an examination of the consistency of the observed rankings reveals.

These not-too-restrictive assumptions are sufficient to generate data-sets for the Monte Carlo experiment. Three data-sets comprising a different number of latent value dimensions are generated. First, I will create a data-set with a one-dimensional structure, which will then be extended by additional dimensions. Since the existing dimensions remain unchanged, when a further dimension is added, the three data-sets have a nested structure. When the measurement model under investigation is used to analyse the artificial data, its capacity to assess the dimensionality correctly can be clarified. Provided that the model meets the expectations and that the fit statistics are not seriously affected by the skewed distributions, the model fit should be sufficient, as soon as all dimensions built into the data-sets have been extracted.

The generation of the data-sets starts with the creation of three uncorrelated, normally distributed random variables – each representing one independent value dimension – with a zero mean and a standard deviation of one.[10] The number of cases is set at 700, that is, roughly the average of the three samples analysed in the previous section. In order to calculate the d_{ij} evaluations, it is necessary to define the positions of the items on the value dimensions. By analogy with Inglehart's measurement instrument, twelve items are provisionally distributed on the dimensions, so that their positions roughly reflect the pattern or observed factor loadings (Table 3). Consequently, when the measurement model is applied to the artificial data, a similar factor pattern should be extracted from the artificial data. Once that the item-positions on the dimensions are defined, the Euclidean distances can be calculated and transformed into rankings. As in the original measurement procedure, the items are ranked in two separate groups of four and eight items.[11]

Yet, as mentioned above, the artificial data should possibly have the same properties as the observed rankings. This is especially important regarding the distributional properties of the rankings, which hinder an unequivocal assessment of dimensionality. The reason is that the skewed distributions of the rankings may distort the χ^2-test, which is crucial when deciding on the number of underlying dimensions. In principle, the artificial rankings become more skewed the more the respective item is moved towards the extremes of a value dimension. Hence, the distributional properties of the artificial data can be adjusted to the observed rankings by varying the item-positions.

As a reference, I choose the German sample, where the deviations from normal distribution are the most pronounced. Table 4 presents the respective descriptive statistics for the original German data, and the artificial one-dimensional data-set. As in the previous section, one item of each separate ranking has to be excluded from analysis, in order to correct for ipsativity. It turns out that the observed distribution can be very well reproduced by moving the items on the dimensions. The same also holds for the two- and three-dimensional artificial data-sets (not shown).

Artificial and observed rankings should not only have similar distributions, the correlations between the rankings should also be similarly strong. The level of correlation among the artificial rankings is primarily a function of the amount of error variance (ε). It turns out that the level of correlation among the artificial ranks approaches the observed level, when the variance of the error term is set at 1.5. For example, the resulting correlations range from -0.49 to 0.37 for the one-dimensional data-set, which is almost the same as the correlation observed in Germany (-0.49 to 0.34). A very similar range is found for the other data-sets, observed as well as artificial. Now, the artificial rankings with known dimensionality are ready for analysis. This allows us to test the capacity of the measurement model and fit statistics to filter out the known underlying structure.

The analysis follows the same steps as in the previous section. First, the number of underlying dimensions is determined. Then, the structure of the dimensions is inspected. The factors are extracted step by step with ULS, and after each extraction the model fit is evaluated. An overview of the results is found in Table 5. The main conclusion is that, despite the violated distributional assumptions, the χ^2-test is unbiased. The fit is not acceptable, until all built-in dimensions have been extracted; for the correct models, in contrast, it is very good. It is noteworthy that the behaviour, not only of the χ^2-test but also of the GFI is almost the same as in the empirical application (cf. Table 2).

Table 4. *Monte Carlo simulation: distributions of observed versus artificial rankings*

Descriptive statistics	Observed rankings (Germany, N = 822) Mean	Standard deviation	Skewness	Kurtosis	Artificial rankings (N = 700) Mean	Standard deviation	Skewness	Kurtosis
Item 1	1.9	1.0	0.6	−0.6	2.1	1.0	0.4	−1.0
Item 2	2.9	1.0	−0.6	−0.7	2.7	1.0	−0.4	−0.9
Item 3	1.9	0.9	0.7	−0.4	2.0	1.0	0.6	−0.7
Item 5	4.4	2.1	0.0	−1.0	4.6	2.0	0.0	−1.1
Item 6	5.7	1.9	−0.4	−0.8	5.7	2.0	−0.4	−1.0
Item 7	4.6	1.9	0.0	−0.8	4.2	2.0	0.0	−1.0
Item 9	1.9	1.4	1.9	3.9	1.9	1.4	1.7	2.8
Item 10	3.1	1.7	0.8	0.1	3.3	2.0	0.6	−0.7
Item 11	5.3	2.1	−0.4	−0.9	5.0	2.0	−0.2	−1.0
Item 12	5.6	2.0	−0.7	−0.5	5.5	1.9	−0.4	−0.6

Table 5. *Monte Carlo simulation: assessing the number of latent dimensions*

		'True' structure of the data-set N = 700								
Extracted		one-dimensional			two-dimensional			three-dimensional		
factors	df[a]	χ^2	P	GFI	χ^2	P	GFI	χ^2	P	GFI
1	33	38.91	0.221	0.999	306.4	0.000	0.962	204.7	0.000	0.978
2	33 (23)	–	–	–	28.2	0.704	0.999	85.6	0.000	0.994
3	33 (13)	–	–	–	–	–	–	27.0	0.760	0.998

[a]Degrees of freedom for the respective extraction step and for the complete model (in brackets).

Given that the results of Monte Carlo studies strongly depend on the defined simulation conditions, one may ask how robust the finding is. Therefore, I have attempted to test the influence of some crucial conditions on the reliability of the χ^2-test. Sample size, error variance, and item discriminability are considered as relevant conditions. Regarding sample size, it can be shown that the test is robust. Even when size of the artificial samples is increased by factor two or three, only the correct models are accepted. When the influence of measurement error is investigated, the test again proves to be very reliable, though only above a certain level of random error. As Figure 2 shows, the correct one-dimensional model is accepted if the variance of the error term is greater than one ($p > 0.05$). In contrast, the true model would be rejected at lower error levels. This finding, which is perhaps rather astonishing at first sight, seems to result from the increase in correlations between the rankings accompanying every error reduction. Above a certain correlation level, the absolute size of the residuals begins to exceed what might be expected by chance, even though their relative magnitude remains unchanged. In addition, the loss of information implied by the ordinal level of measurement may reduce the fit more, when there is less noise in the data. However, the fit statistic gives accurate results, as soon as the correlations are constrained to the empirically observed range by

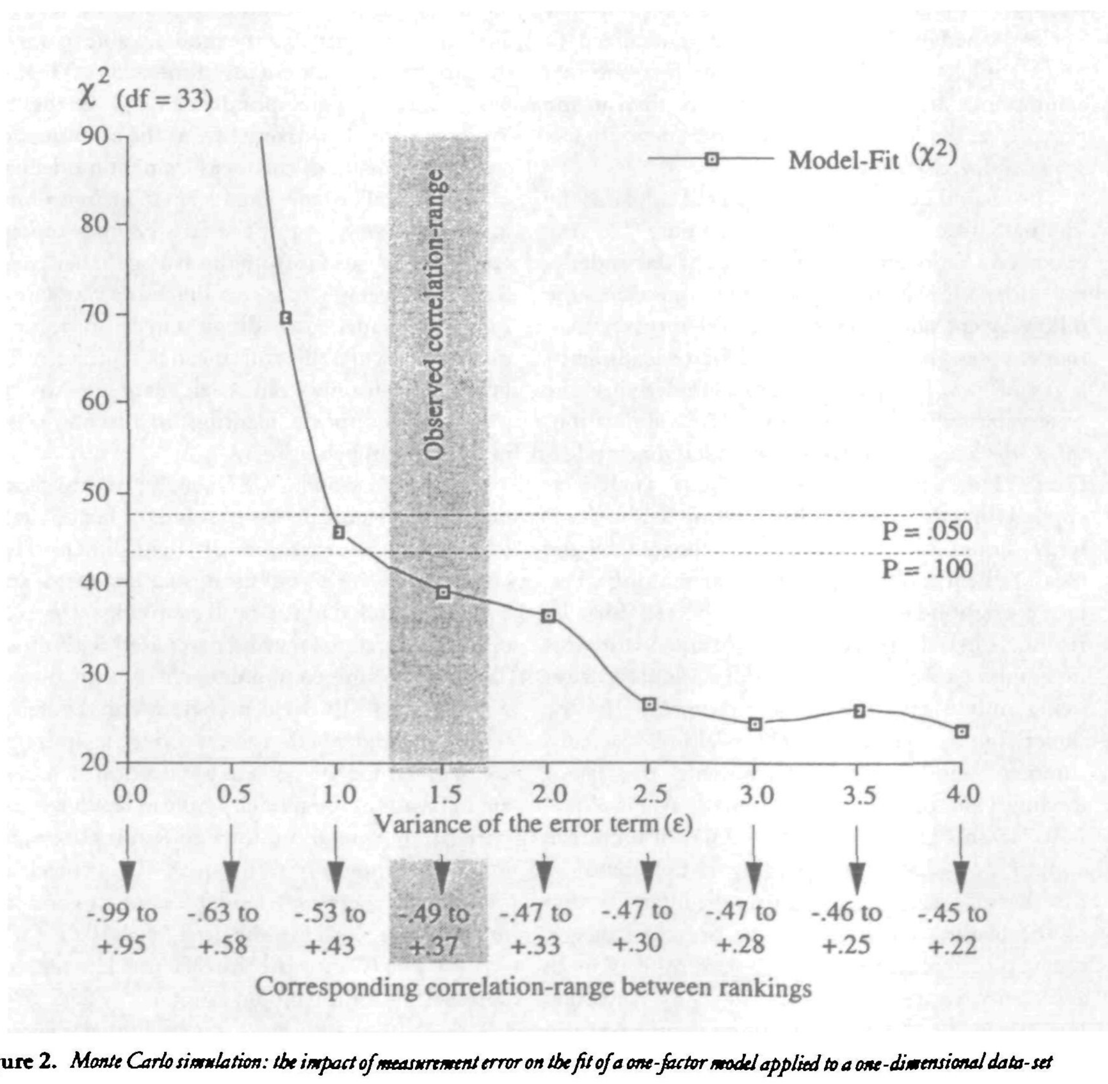

Figure 2. *Monte Carlo simulation: the impact of measurement error on the fit of a one-factor model applied to a one-dimensional data-set*

blowing up the error term slightly. This is true for the grey area in Figure 2.

A third condition that may influence model fit is item discriminability, that is, the degree of distinction between the item meanings with respect to a latent dimension. In fact, it can be demonstrated that the number of factors may be underestimated, when all items are closely clustered on a dimension. In this case, a sufficient fit may result, even if a minor factor with small variance has not been extracted. Much more important, however, is the fact that – even for large samples and a very unreliable measurement – the number of dimensions is not overestimated, that is, the extracted factors are not artefacts. Since the present simulation draws strongly on the Inglehart items, it remains somewhat unclear whether this result can be generalized to other applications. As long as the distributional assumptions are not violated anymore than in the present example, a generalization nevertheless seems quite reasonable.

The second question to be answered concerns the internal structure of the dimensions. Do the extracted factors give a realistic idea of the underlying dimensions? To answer this question, the following consideration is useful: the measurement model is designed to estimate the factor loadings of a set of non-ipsative evaluations underlying the observable rankings. In the Monte Carlo simulation, the d_{ij} distances represent these underlying evaluations. Thus, when conventional factor analysis is applied directly to the d_{ij}, the resulting loading patterns should be close to those estimated by the measurement model on the basis of rankings. Perfect correspondence cannot be expected for two reasons. First, the transformation of the d_{ij} distances into rankings implies a loss of information; the latter being only a raw, ordinal approximation of the underlying d_{ij}. Second, strictly speaking, the measurement model does not estimate the factor loadings, but their deviations from the average loading (Jackson and Alwin, 1980: 231). In empirical applications, the average loading is unknown, so that there is a certain ambiguity in the interpretation of the loading patterns. For the present study, it seems plausible that the average loading is close to zero, since the items have opposite signs. Nevertheless, the loading pattern is to some degree open to interpretation. In the simulation, the underlying 'true' factor pattern may be calculated by applying conventional factor analysis directly to the d_{ij} distances.[12] The calculation is based on the three-dimensional data-set.

The loading patterns of the underlying non-ipsative factors, as well as those estimated with the measurement model, are presented in Table 6. At first, it is striking how well the structure of the observed value dimensions can be reproduced artificially (cf. Table 3). Of course, the aim was to reconstruct the structure of the observed dimensions by positioning the items, but nevertheless the quality of the reproduction comes as a surprise. The findings are thus in favour of both the design of the Monte Carlo experiment and the measurement model under investigation. Second, the results in Table 6 also suggest that the model is able to uncover the structure of underlying dimensions. There is a high degree of correspondence between the 'true' loadings directly extracted from the d_{ij} evaluations, and the estimates of the measurement model based on rankings. The latter thus seem to be interpretable in the usual way, despite the fact that they represent estimated deviations from the average 'true' loading. Since the average 'true' loadings are very close to zero, the estimated loadings remain more or less unchanged when the 'true' mean is subtracted. This, however, will only hold when there are two item-groups with opposite loadings on a factor, as is the case for the Inglehart items.[13]

One might wonder whether the Alwin–Jackson model is in fact superior to standard factor analysis in capturing the structure of latent dimensions. In order to answer this question, standard factor analysis is also applied to the twelve rankings. Hence, the estimated loadings may be compared with those in Table 6, allowing us to judge whether the measurement model really performs better than the standard technique, and whether the warnings against its use are justified for the present application. The results are in favour of the model. This can be illustrated by correlating 'true' and estimated loadings separately for the competing techniques. The correlations (Pearson's r) between standard estimates and 'true' loadings are 0.92 for the first, only 0.62 for the second, and 0.72 for the third factor. The respective correlations for the model estimates in Table 6 (0.94, 0.87, and 0.96) indicate a much higher degree of correspondence, especially regarding factors two

Table 6. *Monte Carlo simulation: model estimates and 'true' factor loadings*

Loadings	First dimension		Second dimension		Third dimension	
(λ_{ij})	estimated	'true'	estimated	'true'	estimated	'true'
Item 1	0.257	0.639	−0.388	−0.919	0.358	0.377
Item 2	−0.122	−0.483	0.507	0.958	0.064	0.257
Item 3	0.459	0.872	−0.089	0.223	−0.390	−0.640
Item 4	–	−0.975	–	−0.122	–	−0.028
Item 5	1.425	1.062	0.336	0.353	0.137	0.088
Item 6	1.154	0.949	−0.219	0.138	−0.419	−0.500
Item 7	−0.499	−0.814	0.577	0.742	−0.560	−0.763
Item 8	–	−0.983	–	−0.226	–	−0.521
Item 9	0.670	0.472	0.351	0.589	0.459	0.943
Item 10	0.268	0.440	−1.136	−0.943	−0.061	−0.274
Item 11	−1.095	−0.984	−0.102	−0.373	−0.024	−0.093
Item 12	−0.942	−0.955	−0.043	−0.401	−0.148	−0.282

Note: The same items (4, 8) as before are excluded for estimation with the measurement model.

and three. Particularly when more than one factor is extracted, the standard estimate of the loading pattern is biased, and, therefore, the Alwin–Jackson model should be applied.

It is not clear to what extent this conclusion can be generalized. In any event, the effect of ipsativity – corrected quite successfully by the model – may be expected to be even stronger in many applications. The potential bias is especially large when the measurement is based on one ranking only, and/or the number of items is lower than in the present example. All in all, results suggest that under such conditions conventional factor analysis is not appropriate, especially when the final solution has more than one dimension.

The findings – showing that the statistical model performs well – allow us to return to substantive problems. Since results are promising, the model can now be used to address two unresolved problems, namely the importance of country differences in the loading patterns (next section), and the stability of the dimensions over time (final section).

Cross-National Variation in the Dimensions

In this section the importance of national differences regarding the three dimensions, or more precisely, the loading patterns in Table 3, will be assessed. Is there non-random national variation, and if so, do the dimensions nevertheless reflect variants of a common cleavage structure? If the interpretation in Habermas's framework is valid, there should be a common pattern, reflecting the contradiction of system and lifeworld. Nevertheless, a limited, but substantial, variation between the countries may be expected for two reasons. First, one of the 'colonizing' subsystems, namely, the state, is a national institution by definition. Although the autonomy of government seems to be more and more reduced due to the globalization dynamic, particularly of the economy, differences in political systems and programmes can still not be disregarded. Second, actors interpret the basic conflict between system and lifeworld in the light of country-specific traditions of political culture. Thus, their understanding of the conflict will be influenced by specific ideologies and traditions. These arguments suggest that despite an invariant basic structure, the dimensions are unlikely to be identical.

The importance of country variations can be assessed with a multi-sample analysis, where the factor loadings are constrained to be equal over countries (Jöreskog and Sörbom, 1989: 227). The resulting fit should not be significantly worse than that of an unrestricted model, provided that the loading patterns are equal in all countries. The respective models with one, two and three factors are again estimated step by step with ULS. Results show that the loading patterns are not identical for the three countries under investigation (Table 7). It should be noted that the extraction of additional factors is based on the better-fitting models without equality constraints. In addition, since no third factor was found for Germany, the three-dimensional models are compared for the United States and the Netherlands only.

Differences between constrained and unconstrained models in the right column of Table 7 are significant in all cases (Long, 1983: 65 f.). Thus, the loading pattern differs over countries for all three dimensions. The level of cross-national variation is quite low, however, and the factors therefore seem to represent national variants of a common structure, rather than completely different dimensions. This becomes obvious, when the χ^2-gain due to the extraction of a specific dimension is compared with the loss caused by the equality constraints.[14] These gains are approximately six times higher than the respective losses for dimensions two and three, and some fifteen times higher for the first dimension. With respect to dimensions one and two, the finding thus confirms the impression derived from a visual inspection of the loading patterns in Table 3: all loadings have the same signs and, with some significant exceptions, also similar magnitudes for all countries. In contrast, it comes as a surprise that country variation also appears to be relatively small for the third factor, even though this is partly due to the lower power of this test.[15]

Similar evidence for a broader pool of industrial countries, also based on visual comparison of loading patterns, is thus corroborated by this more systematic test (cf. Sacchi, 1994). This is in line with the expectation that the factors tap essentially identical conflict dimensions in different industrial societies. Therefore, any explanation should refer to general characteristics of those societies.

Aggregate and Individual-Level Stability of the Dimensions

The question of whether the dimensions are stable over time is crucial not only in Inglehart's framework, but also for the interpretation suggested above. First, with regard to Inglehart's theory, one of the key assumptions is the so-called 'socialization hypothesis', maintaining that (postmaterialist) value orientations remain stable in adulthood. Without this basic assumption, the central expectations of an intergenerational value change towards a more postmaterialist society cannot be justified. Because of its crucial importance, the assumption has often been tested empirically, with controversial results (van Deth, 1983*a*, 1989; Böltken and Jagodzinski, 1985; Inglehart, 1985, 1989; Langeheine, 1987*a*, 1987*b*; Jagodzinski, 1987*a*, 1987*b*; Mnich, 1989; Sacchi, 1991*b*). The most convincing attempt is that by de Graaf *et al.* (1989), to which I have already referred several times. Their findings suggest that postmaterialist values are very stable in the United States and the Netherlands, and somewhat less stable in Germany. Overall, Inglehart's socialization hypothesis is corroborated.

Building on Habermas, it is difficult to derive precise expectations regarding the stability of postmaterialism. The same also holds for the additional dimensions. However, some guesses may be derived from Habermas's argument that the probability of lifeworld reactions is smaller for the core strata directly involved in the production process

Table 7. *Assessment of country differences in the loading patterns*

	Equality constraints over countries[a]						
Extracted	unconstrained loadings		equal loadings		Country Differences		
factors	χ^2	df[b]	χ^2	df[b]	χ^2-Diff.	df	P
1	599.0	99	675.1	119	76.1	20	0.000
2	230.3	99 (89)	284.6	119 (109)	54.3	20	0.000
3	72.8	66 (46)	95.8	76 (56)	21.0	10	0.020

[a]Extraction builds on the better fitting models without equality constraints.
[b]Degrees of freedom for the respective extraction step and the whole model (in brackets).

(Habermas, 1987: ii. 577; see also de Graaf *et al.*, 1989; Bornschier and Keller, 1984; Sacchi, 1991*a*, 1994). Since the degree of integration into this productive core, that is, into the 'colonizing' subsystems, varies over the life-course, some change may be anticipated at the individual level. When this general argument is combined with empirical findings, even more precise expectations can be derived. It is well known that postmaterialist values are more widespread among young people, especially the well-educated entering late into the labour force. Moreover, the conservative potential captured by the second dimension seems to be strongest among elderly and retired people (Sacchi, 1991*a*, 1994; see also Bornschier and Keller, 1994). Hence, the transitions into and out of the labour force, or the core strata, may be expected to have specific effects on each value-dimension, implying that they are not entirely stable at the individual level. An additional hypothesis concerns their stability at the aggregate level. If the factors indeed capture deep-rooted conflict dimensions emerging from long-term modernization processes, they will be more or less invariant in the short run. Consequently, at the level of countries, the structure of the dimensions should be approximately stable over a medium-range observation period.

In order to test the stability of the dimensions at the aggregate as well as the individual level, the data from both panel waves are analysed simultaneously. The second wave was carried out between 1979 and 1981, that is, about six to seven years after the first wave in all countries (Zentralarchiv für empirische Sozialforschung, no date). Thus, the observation period is long enough to allow for a quite powerful test of the stability assumptions outlined above. In addition, due to the length of the interval, test–retest effects can almost be excluded.

A two-dimensional example of the modified measurement model for the panel is sketched in Figure 3. The submodels for the two waves now analysed simultaneously correspond to the model shown in more detail above (Figure 1). Together with the parameters of these submodels, the model estimates the stability of the latent dimensions over the observation period. The standardized coefficients (β_i) measure the individual-level stability of the dimension, taking into account attenuation. First, an unconstrained baseline model is estimated. Then, it is compared with a model with equal factor loadings for both waves. If the dimensions are stable at the aggregate level, the fit of this second model should not be significantly worse. The three dimensions and the respective stability coefficients are again

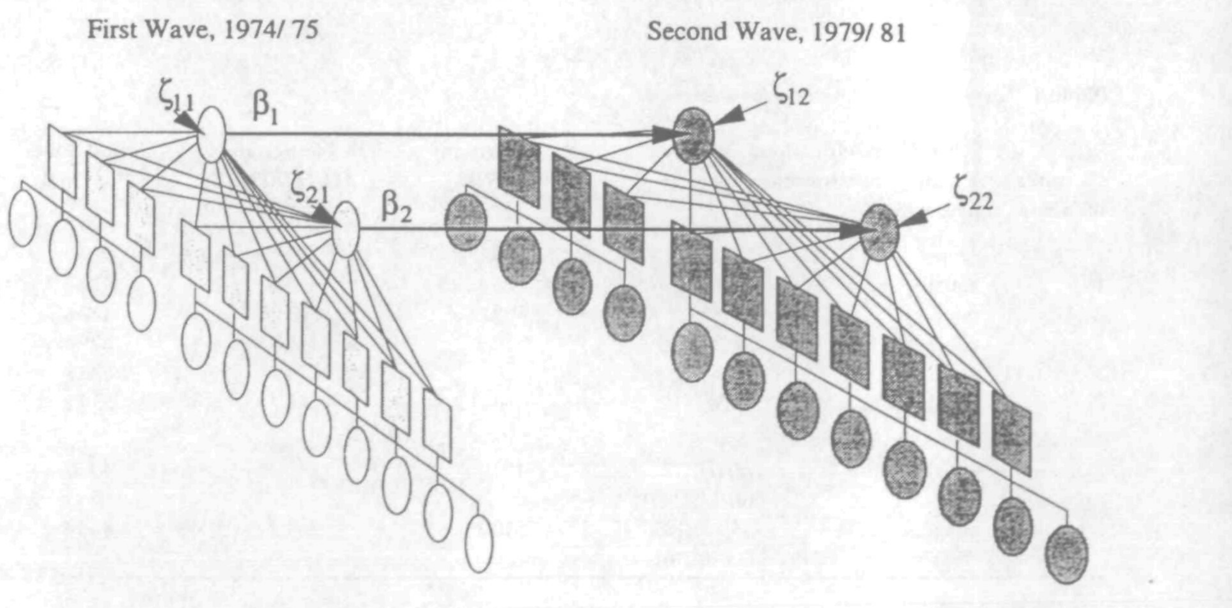

Figure 3. *Two-dimensional measurement model for two panel waves.*

extracted step by step, and all parameters related to a given dimension are fixed on their estimates, when additional factors are extracted. Once all factors are extracted the model fit can be judged and additional restrictions may be introduced, or relaxed, if necessary. This especially refers to the rather strong implicit assumption that the latent dimensions are the only source of correlations between rankings over time. Since the number of dimensions is known from the cross-sectional part of the analysis, it seems reasonable to test this assumption after terminating extraction.

The first extraction step corresponds to a replication of de Graaf *et al.*'s (1989) analysis of the intragenerational stability of postmaterialism. The only difference between their analysis and the present study is that the latter relies on ULS, whereas de Graaf *et al.* used ML. Regarding the first dimension, results of both estimation techniques are similar, although not identical. A first common result indicates that the constrained model with equal loadings for both time-points has a better fit than the unrestricted model (de Graaf *et al.*, 1989: 190, table 8). The χ^2-values show only a very small increase, which is not significant ($p > 0.10$). Thus, the first dimension was stable over the observation period in all countries. Second, the findings are also very similar with respect to the stability coefficients (β_1) for the United States and the Netherlands. Stability is highest for the Dutch sample, and, as already noted by de Graaf *et al.*, the fit even improves when perfect stability is assumed ($\beta_1 = 1$) by setting the residual variance (ζ_{12}) to zero. Possibly, some readers will not have much confidence in stability coefficients of this magnitude, however, the arguments against LISREL's correction for attenuation do not hold in the present case (see de Graaf *et al.*, 1989).[16] Moreover, stability seems to be exceptionally high in the Netherlands even over a panel interval of eleven years. De Graaf *et al.* (1989: 183) found a stability coefficient as high as 0.87 for this time-span. Stability is also very high for the United States, where the respective coefficient of 0.86 comes close to the estimate by de Graaf *et al.* (0.81). The only remarkable difference between the findings concerns individual-level stability in Germany. De Graaf *et al.* – using ML-estimation – found a medium stability coefficient of 0.68. In contrast, the ULS estimate for Germany is 0.48, indicating rather low stability. Thus, both studies lead to the conclusion that stability is lower in Germany than in the other countries. The size of this difference seems to be considerably underestimated when using ML instead of ULS, which is superior, since does it not require a multinormal distribution. As a consequence, the level of stability in Germany is hardly consistent with Inglehart's socialization hypothesis. De Graaf *et al.* (1989: 193) explain the low stability

Table 8. *Comparison of the fit of various two-wave models*

Number of factors	Model specification: Equality constraints over waves[a]	df[b]	West Germany (N = 702) χ^2	The Netherlands (N = 508) χ^2	United States (N = 716) χ^2
1	none	165	550.1	829.7	929.2
1	equal λ_y	174	563.4	830.0	932.3
2	none	165 (154)	308.0	598.2	622.6
2	equal λ_y	174 (163)	322.5	611.0	629.4
3	none	165 (143)	–	334.4	526.9
3	equal λ_y	174 (152)	–	343.9	530.2
Models with item-specific factors:					
2/3/3	equal λ_y, ufl	172/171/171 (161/149/149)	293.4	243.2	405.7
2/3/3	equal λ_y, ufl, ψ	184/183/183 (173/161/161)	330.2	254.2	445.4

[a]The loadings of previously extracted factors are constrained to be equal over waves for *all* models.
[b]Degrees of freedom for the respective extraction step and for the entire model (in brackets).

with an auxiliary argument, namely, the severity of the war time experience leading to an extreme value change. Yet, since the stability seems to be in fact much lower, Inglehart's key assumption is called into question more fundamentally. Nevertheless, de Graaf *et al.*'s argument may be helpful, when postmaterialist values are interpreted in Habermas's sense. Postmaterialist potentials are then rooted in leftist and democratic political traditions, which were radically interrupted during fascism. Later, these political traditions have been renewed only gradually, which is reflected in a relatively low continuity of the respective political values.

Before extracting the second factor, the estimated parameters for the first have to be fixed in order to identify the two-dimensional model. Extraction is based on the best-fitting models with invariant loadings. As in the cross-sectional part of the analysis, the additional factor strongly improves the model fit. The χ^2-statistics decrease by several hundred points, compared to only eleven additional degrees of freedom. When the models with time-specific and invariant loadings are compared, the latter is clearly superior. The increase of the χ^2-values caused by the equality constraints is small and not significant ($p > 0.10$). Thus, the expectation that the second dimension is stable at the aggregate level is fully corroborated. Regarding the Dutch sample, it should be noted, however, that the second factor is equivalent to the third one of the cross-sectional analysis, which has been interpreted as a 'disorientation'-dimension (Table 3).[17] In turn, the factor reflecting the conservative potential is extracted last. As before, the extraction of the third factor builds on the superior models with equal loading.[18] And again, the fit improves considerably when the last dimension is incorporated, especially for the Netherlands. In addition, the expectation of stable loading patterns is also confirmed with respect to the last dimension ($p > 0.10$). Without any exception, the loading patterns from Table 3 are stable in the observation period. This is quite a remarkable finding, since strict metric invariance of factor patterns is found only rarely in empirical research (Cunningham, 1991). Results therefore strongly support the hypothesis that the factors reflect stable political conflict or value dimensions.

The fit of the two-and three-dimensional models is not satisfactory, however, and therefore they should not be accepted without modification. A detailed inspection of the residuals reveals that the models fail to explain the autocorrelations of some items. Probably, this is due to item-specific meanings which are not completely captured by the common dimensions. In order to take such 'memory effects' of items into account, de Graaf *et al.* (1989: 190) incorporated item-specific factors into their model. These factors represent additional latent variables that are exclusively correlated with both rankings of a particular item. When the multidimensional models are extended with an item-specific factor for every item, it turns out that either the estimates of some error variances become negative, or the iteration process does not converge. As Jöreskog points out, negative variances in general emerge from an overparametrization of models (Jöreskog and Sorbom, 1989: 215). This can be avoided by bringing the specific factors into the model step by step, beginning with those items that are responsible for the largest residuals. The respective unique factor loadings (ufl) are constrained to be equal.

It turns out that it is sufficient to allow for just a few item-specific factors in order to obtain an acceptable fit. Two item-specific factors are necessary in Germany (items 1 and 3), three in the Netherlands (items 3, 5, and 6) and the United States (items 1, 3, and 10). Compared to the required degrees of freedom, the extension of the models results in a substantial improvement of fit, especially for the United States and the Netherlands (Table 8). For the one-dimensional model, de Graaf *et al.* found a much larger number of significant item-specific factors. Hence, most of the autocorrelation captured by these factors is now explained by the additional common dimensions.

A final modification of the models concerns the error variances caught in the phantom variable (ψ). For the models at the bottom of Table 8, these variances are constrained to be equal over waves. Results indicate that error variances are invariant for the Netherlands ($p > 0.10$), but not for Germany, nor for the United States ($p < 0.001$). In summary, the models with equal loadings and item-specific factors for particular items perform well in all countries.[19] In addition, the model is slightly improved for the Netherlands, when equal error variances are assumed.

But how about the stability of the additional dimensions at the individual level? As mentioned above, it is difficult to formulate precise theoretical expectations. From the perspective of the 'colonization' thesis, the factors tap collective convictions of social movements, or subcultures. Of course, the degree of integration in the respective lifeworlds as well as the convictions themselves may change during the life-course. Therefore, it is an empirical question as to how much they change, although minimal continuity may certainly be expected. In fact, the stability coefficients of the best-fitting models indicate considerable stability of the captured value orientations for the additional dimensions too (Table 9). For the factor catching the conservative potential, the standardized coefficients are 0.71 for Germany, 0.91 for the Netherlands, and 0.47 for the United States. In Germany, the conservative potential thus seems to be much more stable then the postmaterialist potential. Since conservative traditions were less affected by fascism, this finding is in line with the suggested explanation of the lower stability of postmaterialism in this country. Stability is again very high for the Netherlands – a result pointing to strong political traditions also on side of the conservative potential. In contrast, for the United States the 'potential for withdrawal' seems relatively fluid in the observation period. Maybe this reflects a strengthening of the conservative potential, which became manifest when Reagan was elected as president.

Regarding the 'disorientation'-dimension, the estimated stability coefficients are slightly greater than one for both the United States and the Netherlands, as long as the residual variances are not fixed at zero (Table 9). The finding, in principle indicating perfect individual-level stability, should be interpreted with caution. When the measurement of a given factor has low reliability, LISREL strongly overestimates stability (Jagodzinski, 1984). Regarding the apathy factor, this could be the case, since the standard errors of the loadings are quite large.[20] It seems that the stability of this dimension cannot be accurately assessed with the data-sets analysed in the present study. One might perhaps conclude that the dimension should not be interpreted at all, when its reliability is that low. There are strong arguments against this notion, however. First, the Monte Carlo simulation has revealed that the measurement model does not overestimate the number of dimensions underlying the rankings. Second, the factor can be replicated in many industrial countries, and for different time points. Hence, it is not a random factor. Third, despite the low reliability of the measurement, the factor shows stable and theoretically meaningful correlations with a number of validation criteria (Sacchi, 1994). Thus, at the level of country aggregates, the dimension may be interpreted, although individual-level change cannot be measured reliably enough.

Conclusions

Two major conclusions may be drawn from the present analysis. The first, methodological conclusion concerns the measurement of value orientations. The Monte Carlo experiment reveals that the model developed by Alwin and Jackson is appropriate for factor analysis of ranked Inglehart items. Although results from simulation studies should be generalized with caution, it seems plausible that the model is appropriate not only for the item set under investigation, but for rankings in a more general sense. The finding that the structure of the underlying dimensions can be estimated more accurately by the measurement model than by conventional factor analysis is especially important

Table 9. *Stability coefficients of the best-fitting models*

Standardized coefficients (β)	West Germany	The Netherlands	United States
Potential for resistance	0.48	1.00[a]	0.86
Potential for withdrawal	0.73	0.91	0.47
Disorientation	–	[1.00][a]	[1.00][a]

[a] ζ_{Ω} is constrained to zero when estimating these coefficients.

in this respect. In addition, the determination of dimensionality, that is, of the number of relevant factors, does not present special problems in the context of the model. When these findings are corroborated by further research, the objections raised against ranked preference data founded on the limited possibilities for statistical analysis are obsolete. Hence, there is no methodological reason to abstain from the application of ranking techniques, particularly when measuring value orientations, or, in general, when such techniques are superior for theoretical reasons.

The second, substantive conclusion concerns the dimensionality problem. Results indicate that the Inglehart items tap two or even three independent value or conflict dimensions. The structure of these dimensions is similar for all countries under investigation; furthermore, it remains unchanged over an observation span of approximately six years. Moreover, the captured values show considerable stability at the individual level, although substantial country variation is observed in this respect. The findings are in line with the interpretation referring to Habermas's thesis of a 'colonization' of the lifeworld suggested by previous research. According to this explanation, the dimensions catch different political potentials which react to the subordination of the lifeworld by the econo-administrative complex. In particular, the dimensions are attributed to a postmaterialist, egalitarian potential of the 'new' left, a conservative, authoritarian potential for law and order, and an anomic, lower-class potential left behind by the disintegrating traditional 'old' left.

This notion has been further substantiated by the present analysis in two respects. First, all dimensions prove to be stable in a medium range. Taking into account the fact that the captured values or potentials have a strong impact on various forms of political attitudes and activities, the notion that the dimensions reflect fundamental political cleavages of industrial societies is supported. Second, country variation in the dimensions is small, although not random. This result confirms the supposition that they reflect a political cleavage structure common to advanced industrial societies. All in all, Habermas's theory seems to provide a promising basis not only for a reinterpretation of postmaterialism, but also for research on political values and conflicts in general.

Notes

1. For example, after applying and 'unfolding' technique van Deth (1983*b*: 418) concludes: 'The only firm, but very important, conclusion that this analysis produces is that there is no single joint scale underlying the orderings of all, or most, of the items.'
2. Besides Inglehart's value concept, those of Allport (1960), Kohn (1969), Lenski (1961), and Rokeach (1973) are widely recognized examples.
3. A formal derivation of the model is part of the cited articles.
4. A general tendency to give positive answers (acquiescence) seems to be the reason that the postmaterialist dimension is split on two independent factors, when rating is applied. The two extracted factors oppose respondents giving high scores to all items to those preferring materialist, and postmaterialist items, respectively. When ratings are subsequently transformed into ranks, however, the well-known postmaterialist dimension reappears (cf. Bacher, 1987). Thus, apart from the disturbing preference for positive answers inherent in ratings, both measurement strategies produce similar results.
5. The expected effect size refers to an unstandardized coefficient in the ordinal metric of rank-indicators. The resulting λ_y-matrix for a one-dimensional model for the Inglehart items is presented as an appendix (see also de Graaf *et al.*, 1989).
6. Because the ranking of eight items is incomplete, the items on positions four and five are both coded as 4.5.
7. Assuming that the deviations from multinormal distribution are negligible, standardized residuals can be interpreted as standard normal deviates. Thus, in large samples residuals with absolute values exceeding 2.58 are in indication of specification errors (Jöreskög and Sörbom, 1989: 32). For the accepted models, however, only the residual for the covariance of items 3 and 6 in the US sample slightly exceeds this critical value.
8. The reason is that the loading estimates of the model represent deviations from the mean loading of the original non-ipsative factor structure (Alwin and Jackson, 1981: 322; de Graaf, 1988: 50).
9. This objection was made by an anonymous reviewer.
10. The artificial data is generated by SPSS.
11. The coding of the rank indicators is the same as in previous section.
12. As in the previous section, the factors – or more precisely, principal components – are extracted step by step with ULS, now including all twelve

items, however (Jöreskog and Sörbom, 1986: iii. 102).

13. The signs of the loadings are the same for all three factors, when standard factor analysis – giving a direct, although potentially biased estimate – is applied (Sacchi, 1991*a*).
14. The overall χ^2-gain can be calculated by summing up the nation-specific statistics in Table 2, and then subtracting the resulting sums of subsequent models.
15. Standard errors of the loadings are largest for the third factor.
16. Jagodzinski (1984) has shown that LISREL tends to overestimate stability coefficients, when the reliability of the measurement is very low. However, since the zero-order correlations between the rank indicators are substantial and the standard errors of all loadings are small, his argument is not valid with respect to the first dimension, as de Graaf *et al.* (1989: 194) have noted already. Moreover, two possible sources of additional autocorrelation, namely test–retest effects and 'memory' effects of particular items, can be excluded in the present application. The former are unlikely to occur due to the length of the panel-interval, and the latter have been controlled in the accepted models.
17. Probably, the 'disorientation'-dimension has a stronger impact on the rankings of the second wave, and, therefore, the respective factor is extracted first when both waves are analysed simultaneously. This conclusion is suggested by the model parameters, showing that the variance of the factor has increased over time.
18. For the Netherlands, the residual variance (ζ_{22}) is constrained to zero.
19. Inspection of the residuals reveals that the fit of these models is acceptable: standardized residuals form a straight line hiding the diagonal of the respective plot. The low probability levels of the accepted models ($p = 0.000$) thus seem to be a consequence of sample size (see Jöreskog and Sörbom, 1989: 43).
20. Only some of the items have significant loadings; however, standard errors and significance tests may be biased in the present application.

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Appendix

The $\lambda_{i,1}$-parameters in the matrix below represent the factor loadings on the first factor (η_1 in Figure 1), and the remaining columns the phantom variables (η_{3-14} in Figure 1). Except the loadings, all

$$
\begin{array}{l|rrrr|rrrrrrrr}
\lambda_{1,1} & -0.250 & 0.750 & -0.250 & -0.250 & & & & & & & & \\
\lambda_{2,1} & -0.250 & -0.250 & 0.750 & -0.250 & & & & 0 & & & & \\
\lambda_{3,1} & -0.250 & -0.250 & -0.250 & 0.750 & & & & & & & & \\
\hline
\lambda_{4,1} & & & & & 0.875 & -0.125 & -0.125 & -0.125 & -0.125 & -0.125 & -0.125 & -0.125 \\
\lambda_{5,1} & & & & & -0.125 & 0.875 & -0.125 & -0.125 & -0.125 & -0.125 & -0.125 & -0.125 \\
\lambda_{6,1} & & & & & -0.125 & -0.125 & 0.875 & -0.125 & -0.125 & -0.125 & -0.125 & -0.125 \\
\lambda_{7,1} & & & 0 & & -0.125 & -0.125 & -0.125 & 0.875 & -0.125 & -0.125 & -0.125 & -0.125 \\
\lambda_{8,1} & & & & & -0.125 & -0.125 & -0.125 & -0.125 & 0.875 & -0.125 & -0.125 & -0.125 \\
\lambda_{9,1} & & & & & -0.125 & -0.125 & -0.125 & -0.125 & -0.125 & 0.875 & -0.125 & -0.125 \\
\lambda_{10,1} & & & & & -0.125 & -0.125 & -0.125 & -0.125 & -0.125 & -0.125 & 0.875 & -0.125
\end{array}
$$

parameters are fixed when estimating the model, where the small submatrix at the top corrects for the ipsative properties of the four-item ranking (loading 1 to 3), and the larger submatrix at the bottom for those of the ranking with eight items (loading 4 to 10). Additional factors (η_2 in Figure 1) may be extracted by introducing an additional column of loadings and fixing the loadings of previously extracted factors on the estimates.

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